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Do Migrant Remittances Complement Domestic Investment? New Evidence from Panel Cointegration

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Abstract

This paper examines whether migrant remittances “crowd in” or “crowd out” domestic investment in developing countries. Using recently developed panel cointegration techniques that account for cross-sectional dependence, structural breaks and regime shifts, the paper shows that remittances form a long-run equilibrium relation with domestic investment. The results of the panel vector error correction model reveal the absence of a short-run relationship but the presence of a long-run bidirectional link between remittances and investment. Thus, remittances drive investment while investment itself cause more remittances, suggesting that remittances are not only driven by altruistic motives but also investment motives.

Keywords: Remittances, Investment, Panel Cointegration

JEL classification: F24, E22, C23

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1 Introduction

Over the last few decades, migrant remittances have taken a more prominent role in developing countries. As a result, the question of whether they crowd-in or crowd-out domestic investment has become an important policy issue. In general terms, the macroeconomic effects of remittances largely depend on whether they act as pure compensatory transfers or capital flows (Chami et al. 2005). In the first case, altruistic motives dominate in the sense that the migrant is concerned with the well-being of his/her relatives. In the latter case, though, self-interest dominates, such that the migrant retains some sort of ownership over the assets. In both cases, however, the response of the economy to increases in remittances could be either negative or positive.

On the one hand, remittance flows can have negative effects on the recipient economy through their adverse influences on income distributions (Orrenius et al. 2010), household's labour supply and savings rates (Chami et al. 2005). In addition, similar to any other resource inflow, sustained levels of remittances tend to be associated with "Dutch disease" effects (Amuedo-Dorantes et al. 2005) as well as increases in conspicuous consumption rather than productive investments (Chami et al. 2005).

On the other hand, there is considerable evidence showing that, although remittances may mainly go to consumption, a substantial portion of it goes to human capital formation in the form of better nutrition, schooling and health (Gupta et al. 2009). Moreover, increased consumption and even "unproductive" investments (e.g. real estate) can have significant multiplier effects, encouraging more capital accumulation and growth through spillover effects (Ratha, 2003; Gupta et al. 2009).

Evidence also suggests that remittances tend to reduce households' credit constraints and thus boost the depth of the financial sector (Guilamo and Ruiz-Arranz, 2009; Aggarwal et al. 2011). Furthermore, it has been shown that remittance receiving households, on average, tend to save and invest more than other comparable households (Adams, 2007). Other studies found that remittances are associated with poverty reduction (Adams and Page, 2005) and higher educational attainments (Rapoport and Docquier, 2006). Finally, remittance flows have been found to act more counter-cyclically than other types of inflows and thus are a more stable

source of foreign exchange at times of economic difficulties (Combes and Ebeke, 2011; Chami et al. 2009).

The objective of this study is to contribute to this literature but we depart from the existing literature in a number of ways. First, we use recently developed panel cointegration tests that can handle a number of econometric issues, including cross-sectional heterogeneity, structural breaks and endogeneity concerns. Second, we examine the long-run relationship between remittance inflows and domestic investment. Third, we apply panel error correction methods to uncover the short-run dynamics in the relationship. Finally, we conduct a panel Granger causality analysis in order to establish whether the long and short-run effects are indeed of a causal nature.

The paper is organised as follows. Section 2 sets out the econometric analysis, presenting the techniques used as well as the findings while Section 3 concludes.

2 Empirical analysis

To examine the relationship between remittances and domestic investment, we use a balanced panel of 47 developing and emerging economies over the period 1980-2006¹. The model takes the following form:

$$INV_{it} = \alpha_i + \gamma_{it} + \beta REM_{it} + \varepsilon_{it} \quad (1)$$

where α_i and γ_{it} are, respectively, country specific fixed and time effects, capturing any country-specific unobservables that are relatively stable over time and ε_{it} is the error term. INV_{it} is the share of investment in GDP for countries $i=1, \dots, N$ and time periods $t=1, \dots, T$, and REM_{it} is the share of remittances in GDP, both sourced from World Development Indicators (2011).

As is the standard norm in panel cointegration studies (see for example, Crowder and de Jong, 2011; Herzer and Grimm, 2012), equation (1) is a parsimonious specification that solely focuses on the bivariate long-run link between REM and INV . The validity of this specification,

¹ The sample selection is based on the availability of consistent data.

however, requires that the variables in (1) are nonstationary or, more precisely, integrated of the same order. In that case, they would have a stationary error term, implying that they constitute a cointegrating vector (Asteriou and Hall, 2007). Once a set of variables form a cointegrating relation, such (long-run) relationship should exist even if more variables are added to the model (see for example, Herzer and Grimm, 2012).

2.1 Panel stationarity tests

In estimating equation (1), we first test the time series properties of the variables using the panel unit root tests developed by Levin, Lin and Chu (2002) (*LLC*) and Im, Pesaran and Shin (2003) (*IPS*). The *LLC* is an extension of the standard (Augmented) Dickey-Fuller test and assumes parameter homogeneity while the *IPS* allows for heterogeneity across the panel and serial correlation in the error terms. Both the *LLC* and *IPS* may lead to erroneous results if there is cross-sectional dependence among the panel members emanating from, for example, common effects. Hence, we also report the cross-sectionally augmented panel unit root test (*CIPS*) proposed by Pesaran (2007), which takes into account possible cross-sectional dependence.

Table 1: Panel unit root test results

	LLC statistics		IPS statistics		CIPS statistics	
	Levels	Diff	Levels	Diff	Levels	Diff
INV_{it}	-0.41	-1.22**	-2.21	-3.09**	-2.22	-2.73**
REM_{it}	-0.23	-1.05**	-1.41	-2.81**	-2.16	-2.70**

Notes: The tests are: Levin, Lin and Chu (2002, LLC), Im, Pesaran and Shin (2003, IPS) and Pesaran (2007, CIPS). ** indicates the rejection of the null of non-stationarity at the 5% level or better. Two lags used to account for autocorrelation and the tests include intercept and trend in levels.

Table 1 reports the results of the unit root tests which indicate that we cannot reject the null hypothesis of a unit root in levels, suggesting that the variables are non-stationary. However, the series are stationary in first-differences, implying that they are integrated of order one, $I(1)$. Hence, we can now proceed with panel cointegration tests to explore whether there is a long-run equilibrium relationship between *REM* and *INV*.

2.2 Panel cointegration tests

Having established that the variables under study are $I(1)$, we now explore whether there is a long-run cointegration between *INV* and *REM*. To this end, we implement the residual based

panel cointegration test developed by Kao (1999) which is an ADF-type test. The null hypothesis tested here is that there is no panel cointegration against the alternative of cointegration based on the assumption of homogenous cointegrating vectors. Since the assumption of homogeneity among the cross-sectional units may be too strong, we also report the Pedroni (1999, 2004) panel cointegration test which offers considerable flexibility as it allows for heterogeneity in the long-run cointegrating vectors. Pedroni (1999, 2004) constructs seven test statistics which capture both the within- and between-dimensions of the panel.

However, an important shortcoming with the above panel cointegration tests is that they impose a common factor restriction - that is, they assume that the long-run parameters for the level variables are equal to the short-run parameters of the variables in their first differences. As shown by Westerlund (2007), when this assumption does not hold, the above cointegration methods suffer from a significant loss of power. Therefore, in addition to the above methods, we also report more appropriate panel cointegration tests proposed by Westerlund (2007). Westerlund (2007) sidesteps the assumption of a common factor restriction by utilising the structural (rather than residual) dynamics. The Westerlund test can handle serially correlated residuals, country-specific intercept and slope parameters along with trend terms. Westerlund (2007) develops four different statistics which can be used to establish the existence of a panel cointegration. Two of them are *panel* tests (denoted P_τ and P_α), testing the alternative hypothesis that the panel is cointegrated as a whole ($H_1^p: \alpha_i = \alpha < 0$ for all i). The other two are *group-mean* statistics, (denoted G_τ and G_α), which test the alternative that at least one element in the panel is cointegrated ($H_1^g: \alpha_i < 0$ for at least one i). Thus, the panel tests assume that α_i is homogenous for all i while the group-mean tests do not require this.

To formally examine whether the panel members are indeed independent, we apply the *CD* test proposed by Pesaran (2004). Pesaran (2004) shows that the *CD* test is robust to a single as well as multiple breaks in the slope parameters and/or in the residual variances of the individual regressions.

Given the length of the time period we cover and the heterogeneity of the countries under study, it is highly likely that our variables of interest may have been influenced by various

shocks emanating from, for example, regime and policy changes. Thus, to fully understand the relationship between *INV* and *REM*, structural breaks and regime shifts need to be accounted for. In this study, as an additional robustness, we implement the panel cointegration test proposed by Westerlund and Edgerton (2008), which accounts for both structural breaks and cross-sectional dependence. Westerlund and Edgerton (2008) develop two different tests that allow for unknown structural breaks in both intercept and slope of the cointegrating model, heteroskedastic and serially correlated errors as well as time trends. The location of the structural breaks may be at different dates for the cross-sectional units.

Table 2: Panel cointegration test results

Kao (1999)		T-statistic	Prob.
	ADF	-2.982**	0.001
Pedroni (1999, 2004)			
	<i>Panel ν – stat</i>	-5.235**	0.000
	<i>Panel ρ – stat</i>	-2.808**	0.007
	<i>Panel PP – stat</i>	-6.736**	0.000
	<i>Panel ADF – stat</i>	-8.647**	0.000
	<i>Group ρ – stat</i>	1.087	0.221
	<i>Group PP – stat</i>	-3.585**	0.000
	<i>Group ADF – stat</i>	-5.480**	0.000
Westerlund (2007)			
	<i>Gτ</i>	-2.314**	0.000
	<i>Gα</i>	-7.765**	0.000
	<i>Pτ</i>	-14.221**	0.030
	<i>Pα</i>	-6.588**	0.010
Pesaran (2004)			
	<i>CD statistic</i>	12.660**	0.000
Westerlund and Edgerton (2008)	<i>Model</i>	<i>Z$\tau(N)$</i>	<i>Z$\phi(N)$</i>
	<i>No break</i>	-11.531**	-20.553**
	<i>Level break</i>	-8.352**	-17.851**
	<i>Regime shift</i>	3.700	0.060

Notes: The null hypothesis of the Kao and Pedroni tests is that the variables are not cointegrated and the lag lengths are based on Schwartz Information Criterion with a maximum number of 3 lags. Under the null, the Pedroni tests are distributed as normal and their finite sample distribution are tabulated in Pedroni (2004). For the Westerlund (2007) test, the optimal lag/lead length is determined by Akaike Information Criterion with the maximum of lags set equal to 3 and the width of Bartlett-kernel is set to 3 (bootstrapped robust *p-values* reported). The Pesaran (2004) CD test takes cross-sectional independence as the null and its associated *p-values* are for a one-sided test based on normal distribution. The lag length selection of the Westerlund and Edgerton (2008) test is based on an automatic procedure and 3 breaks are used based on grid search at the minimum of the sum of squared residuals. The *p-values* are for a one-sided test based on the normal distribution. ** denotes significance level at the 5% or better.

In the top panel of Table 2, we report the results of the Kao (1999) test which strongly

rejects the null hypothesis of no cointegration between *INV* and *REM*. The null of no cointegration is also rejected when we allow for heterogeneous cointegrating vectors using the Pedroni (1999, 2004) tests. The table also reports the results based on Westerlund (2007). To account for cross-sectional dependence, bootstrapped robust p -values are reported (based on 500 replications). The results indicate that the null hypothesis of no cointegrating relationship can be rejected irrespective of whether we treat α_i as homogenous (tests P_τ and P_α) or not (tests G_τ and G_α). Thus, there is a strong evidence of a cointegrating relationship between *REM* and *INV*.

To formally establish the existence of a cross-sectional dependence, we apply the *CD* test which strongly rejects the null hypothesis of no cross-sectional dependence (see Table 2). Thus, a failure to take this into consideration may result in biased results.

Finally, we consider the effects of structural breaks and regime shifts on the long-run relationship between *REM* and *INV* using the test developed by Westerlund and Edgerton (2008). Table 2 reports the results for three cases (no-break, level-break and regime-shift). When possible structural breaks are ignored (the no-break case) or accounted for (the level-break case), the null hypothesis of no cointegration can be rejected. However, when we consider regime shifts we fail to reject the null of no cointegration.

To sum up, we find that there is a long-run relationship between *INV* and *REM*. This link is robust to heterogeneity in the long-run cointegrating vectors as well as to cross-sectional dependence and structural breaks. However, it is not robust to regime shifts. With this in mind, we now estimate the nature of this relationship.

2.3 Long-run estimation

Having confirmed the presence of a cointegration, we apply the within-dimension-based dynamic OLS (*WD-DOLS*) estimator developed by Kao and Chiang (2001) to uncover the effect of *REM* on *INV*. To implement the *WD-DOLS* estimator, we consider the following panel model:

$$INV_{it} = \lambda_i + \beta_1 REM_{it} + \varepsilon_{it} \quad (2).$$

Because our data is non-stationary, the *WD-DOLS* estimator addresses issues of serial correlation and endogeneity concerns by augmenting equation (2) with leads and lags of the first differences of the right hand side (endogenous) variable as follows:

$$INV_{it} = \lambda_i + \beta REM_{it} + \sum_{j=-q}^q \Psi_{ij} \Delta REM_{it+j} + v_{it} \quad (3)$$

where Ψ_{ij} are the leads and lags. The *WD-OLS* estimator is superconsistent, under cointegration, producing unbiased estimates of the long-run cointegrating relationship.

Nevertheless, a particular weakness with the *WD-DOLS* estimator is that it assumes that the slope coefficients are homogenous across the cross-sectional units. However, this pooling assumption, if not true, can result in a serious bias in both static and dynamic panels (Asteriou and Hall, 2007). Thus, as a robustness check, we also estimate our model (equation 2) using the between-dimension mean-group DOLS (*MG-DOLS*) estimator for heterogeneous cointegrated panels suggested by Pedroni (2001). This estimator allows the long-run slope coefficients to vary across countries by running separate regressions for each cross-section and then averaging them, $\hat{\beta} = N^{-1} \sum_{i=1}^N \hat{\beta}_i$. Thus, the estimates can be viewed as the mean value of the individual cointegrating vectors. As emphasised by Pesaran and Smith (1995), group-mean estimators generate more consistent estimates, in the presence of heterogeneous cointegrating vectors, than do within-dimension estimators. In addition, the *MG-DOLS* estimator has better small sample properties (Pedroni, 2001).

As highlighted previously, we need to consider the possible issue of cross-sectional dependency. For example, investment rates and remittance flows in our sample of countries may respond to (unobserved) common external shocks (e.g. global business cycles), meaning that they may become correlated across i . Ignoring this interdependence may result in erroneous estimates. A simple way to deal with this type of error dependence is to demean the data over the cross-sectional units so that the cross-section averages of the variables, say $\bar{x}_t = N^{-1} \sum_{i=1}^N x_{it}$ are subtracted from the observations, say x_{it} . This procedure can mitigate the effects of error dependence (Pedroni, 2001; Levin et al. 2002). Thus, we re-estimate the *WD-DOLS* regressions using demeaned data. This simple strategy, while effective, implies that the unobserved external

factors are the *same* across countries. To the extent that countries have different macroeconomic and institutional environments, for example, it is highly likely that their responses and behaviour towards remittances would be different. To this end, we also apply the Common Correlated Effects Mean Group estimator (*CCEMG*) developed by Pesaran (2006). Applying this estimator, one can rewrite the error term in Equation (2) as having a multifactor structure as follows:

$$\varepsilon_{it} = \omega_i' \mathbf{f}_t + \nu_{it} \quad (4)$$

where \mathbf{f}_t is $k \times 1$ vector of unobserved common factors, which may affect the countries with different intensities, and ν_{it} is country-specific error term, assumed to be weakly dependent across the cross-sectional units. The common factors \mathbf{f}_t are allowed to be correlated with the regressors in Equation (2):

$$x_{it} = \eta_i + \xi_i' \mathbf{f}_t + \varepsilon_{it} \quad (5)$$

where x_{it} is each of our regressors, ξ_i is $k \times 1$ vector of factor loadings, and ε_{it} is the error term assumed to be independently distributed of \mathbf{f}_t and ν_{it} .

To take into account the presence of common effects, Pesaran (2006) suggests that one can approximate \mathbf{f}_t by cross-section averages of the dependent and explanatory variables and then run standard panel regressions augmented with these averages. As shown by a number of studies (e.g. Pesaran, 2006; Pesaran and Tosetti, 2011), this *CCEMG* performs well in small samples and can handle the presence of autocorrelation in the residuals and unit roots in the common factors.

As a final robustness check, we apply Breitung's (2005) two-step estimator which, unlike the above methods, can handle dynamic effects. Following Breitung (2005), it can be shown that a cointegrated model has the following Vector Error Correction Model (VECM) representation (in the case of a VAR[1]):

$$\Delta y_{it} = \alpha_i \beta' y_{it-1} + \varepsilon_{it} \quad (6)$$

where ε_{it} is a white noise error with $E(\varepsilon_{it})=0$ and positive definite covariance matrix $\sum_i = E(\varepsilon_{it}\varepsilon_{jt})$. The matrix β' captures the long-run relationship among the variables and is assumed to be the same across i while a_i and \sum_i are short-run parameters which vary across i . In the first step, the country-specific short-run parameters are generated from separate models for each cross-section unit resulting in country-specific cointegration vectors. In the second step, the long-run cointegration matrix β' is estimated using the pooled regression:

$$\hat{q}_{it} = \beta' y_{it-1} + \hat{v}_{it} \quad (7)$$

where \hat{q}_{it} and \hat{v}_{it} are based on the generated short-run parameters a_i and \sum_i . Breitung (2005) and Breitung and Pesaran (2008) show that this estimator has a normal distribution and corrects for endogeneity in the second step.

Table 3: The impact of *REM* on *INV*

<i>Estimator</i>	<i>REM_{it}</i>	<i>N</i>	<i>Obs</i>
WD-DOLS (Kao and Chiang, 2001)	0.431 [4.460]***	47	1269
WD-DOLS (Demeaned data)	0.222 [1.910]**	47	1269
MG-DOLS (Pedroni, 2001)	0.628 [9.380]***	47	1269
CCEMG estimator (Pesaran, 2006)	0.222 [0.981]	47	1269
2-step estimator (Breitung, 2005)	0.302 [6.293]***	47	1269

Notes: *T-statistics* in parenthesis. ** and *** indicate significance at the 5% and 1% levels, respectively. The DOLS regressions are estimated with two leads and two lags. The regressions include unreported fixed effects.

Table 3 contains the results of the estimates of the long-run effects of *REM* on *INV*. The coefficient of *REM* is positive and highly significant at the 1% level. The magnitude of the coefficient ranges between 0.22 and 0.63, implying that, in the long-run, a one percentage point increase in the *REM* to GDP ratio leads to an increase in *INV_{it}* of around 0.22–0.63 percentage points.

2.4 Short-run dynamics and causality tests

Given that the variables are cointegrated, we set up a panel vector error correction model in order to explore whether the relationship between REM and INV is of a causal nature. To this end, following Engle and Granger (1987), we use the following two-step procedure (Pesaran et al. 1999). First, the long-run model specified in equation (2) is estimated in order to obtain its residuals. Second, defining the lagged residuals from equation (2) as the error correction term, the following error correction model is generated:

$$\Delta INV_{it} = \alpha_{1j} + \sum_{k=1}^p \gamma_{11ik} \Delta INV_{it-k} + \sum_{k=1}^p \gamma_{12ik} \Delta REM_{it-k} + \lambda_{1i} \varepsilon_{it-1} + u_{1it}, \quad (8)$$

$$\Delta REM_{it} = \alpha_{2j} + \sum_{k=1}^p \gamma_{21ik} \Delta REM_{it-k} + \sum_{k=1}^p \gamma_{22ik} \Delta INV_{it-k} + \lambda_{2i} \varepsilon_{it-1} + u_{2it}, \quad (9)$$

where Δ is the first-difference operator; p is the optimal lag length determined by standard information criterion. The null hypothesis of no short-run causality can be examined, respectively, based on $H_0: \gamma_{12ik} = 0$ and $H_0: \gamma_{22ik} = 0$ for all ik . In other words, short-run causality can be tested evaluating the statistical significance of the partial F -statistic associated with the corresponding regressor. On the other hand, long-run causality can be tested by the statistical significance of λ_{1i} and λ_{2i} (the error correction terms), respectively, using T-statistics.

Table 4: Short-run dynamics and causality tests

Dependent variable	Source of causality		
	Short-run		Long-run
	ΔINV	ΔREM	ECT
Equation (8) ΔINV	-	1.260 [0.262]	0.458*** [0.000]
Equation (9) ΔREM	1.920 [0.166]	-	0.019*** [0.009]

Notes: Partial F-statistics are reported with respect to short-run changes in the respective regressor. The ECM is the coefficient of the error correction term. *** indicates significance at the 1% level.

The long and short-run Granger causality tests are reported in Table 4. The results show there is no *causal* relationship between REM and INV in the short-run as both respective (lagged) regressors are not significantly different from zero at standard confidence levels. However, in the long-run, we find a significant two-way causal relationship. That is, increases in INV are both *a result of* as well as *a cause of* increases in REM .

2.5 Discussion of the findings

Our central findings show that remittances have a robust long-run effect on domestic investment in developing countries. This result is consistent with the recent findings by Ziesemer (2010), who has shown that remittances enhance fixed capital formation directly as well as indirectly through their beneficial influences on public expenditures on education and literacy. The idea that remittance flows improve human capital (e.g. education, nutrition and health) has been confirmed by a number of studies (see for example, Acosta et al. 2007; Calero et al. 2009). Hence, these flows are likely to have positive effects in the long-run. Our findings are also in line with the results of Nsiah and Fayissa (2011) who found that remittances are positively related to economic development in developing countries. Unlike their study, however, we pay particular attention to the properties of the variables under study as well as the underlying assumptions of the econometric techniques. Given that we employ more superior estimation methods, our results should be more reliable.

Our causality analysis show that there is a bidirectional causal relationship between *REM* and *INV*. This could be because of the multiplier effects generated by the expenditures of remittance-receiving households may be encouraging more investment. Alternatively, it could be that the households themselves may be making small capital investments. In the latter case, this could generate more remittance flows if we assume that the migrant is not just altruistic but also self-interested. In other words, if remittance-receiving households engage in successful business ventures, the migrants may send more remittances in order to enhance their own wealth². Results by Alleyne et al. (2008) confirm that remittances are not only driven by altruistic motives but also investment motives. Thus, remittances may drive investment while investment itself may cause more remittances. These ideas are consistent with the theoretical work by Le (2011), who has shown that remittances can act as a useful source of finance for investment projects particularly when the domestic financial system is sufficiently developed.

² This assumes that the migrant and the remittance-receiving household can overcome issues of adverse selection and moral hazard and that they can trust each other.

3 Concluding Remarks

The objective of this study was to establish whether there is a long-run stable relationship between domestic investment and remittances in developing countries. Using recently developed panel cointegration techniques, we show that there is a long-run relationship between investment and remittances. This result is robust to cross-sectional dependency as well as structural changes. Upon estimating the nature of the long-run equilibrium relationship, the study found that remittances have a statistically significant positive effect on investment in the long-run. Moreover, the Granger causality tests show that the long-run link between remittances and investment is bidirectional and of a causal nature.

The overall findings suggest a number of important policy implications. First, developing countries should improve the effectiveness of remittance inflows. A particular channel is the financial system. Thus, developing countries should develop their financial sectors in order to allow remittance-receiving households to have the facilities needed for productive investments. Given that remittances tend to boost the level of deposits and credit in banking system (Aggarwal et al. 2011), a well-developed financial system would likely generate more benefits. In the same vein, they should adopt policies that may reduce the transaction costs attached to receiving the funds so that households can get their remittances as smoothly as possible. One way to do this is to reduce red tape, but perhaps, more importantly, competition should be encouraged among money transfer companies.

Overall, the important role migrant remittances can play in economic development is not a trivial matter. As shown in this study, remittances can improve the economic performance of developing economies by augmenting the rate of capital accumulation.

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